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Health Reform, Population Policy and Child Nutritional Status in China

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Abstract

This paper examines the determinants of child nutritional status in seven provinces of China during the 1990s, focusing specifically on the role of two areas of public policy, namely health system reforms and the one child policy. The empirical relationship between income and nutritional status, and the extent to which that relationship is mediated by access to quality healthcare and being an only-child, is investigated using ordinary least squares, random effects, fixed effects, and instrumental variables models. In the preferred model – a fixed effects model where income is instrumented – the author find that being an only-child increases height-for-age z-scores by 0.119 of a standard deviation. The magnitude of this effect is found to be largely gender and income neutral. By contrast, access to quality healthcare and income is not found to be significantly associated with improved nutritional status in the preferred model. Data are drawn from four waves of the China Health and Nutrition Survey.

This paper—a product of the Human Development Network, Health, Nutrition and Population Department—is part of a larger effort in the department to understand the effects of socioeconomic status and policy change on health status. Policy Research Working Papers are also posted on the Web at http://econ.worldbank.org. The author may be contacted at cbredenkamp@worldbank.org.

Health Reform, Population Policy and Child Nutritional Status in China

by

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1. Introduction

Over the course of the 1990s, there was a remarkable fall in the national prevalence of child undernutrition in China. According to the 1992 China National Nutrition Survey and the 2002 China National Nutrition and Health Survey, the prevalence of stunting among children under five fell by 42% between 1992 and 2002 (Wang et al. 2007). This was the largest percentage decline in undernutrition experienced by any country during that decade (WHO 2007).

In explaining undernutrition, it is usually observed that child nutritional status is the outcome of a complex interaction of immediate, underlying and basic determinants (Mosley and Chen 1984; UNICEF 1990; Jonsson 1993; Smith and Haddad 2000; Bryce et al. 2008). Empirically, income has been found to be among the most critical determinants of a household's decision to investment in its children's nutritional status (see, for example, the reviews by Behrman and Deolalikar 1988; Strauss and Thomas 1995; Chamarbagwala et al. 2004), but other socioeconomic factors and policies can substantially mitigate or accentuate the income effect.

This paper attempts to shed some light on how the observed relationship between income and child nutritional status in China in the 1990s was mediated by two specific areas of policy. First is the one child family policy to which a sharp increase in the number of only-children has been attributed. Second is a protracted period of health system reform which has resulted in increased variation in the accessibility, cost and quality of healthcare, with potential implications for children's consumption of health services, their incidence of illness and, thus, nutritional status.

The paper commences with a brief description of the one-child policy and a discussion of some key characteristics and consequences of the health system reforms. There follows an elaboration of the theoretical model, a discussion of the data, an explanation of the analytical approach (model specification), and a description of the measurement of key variables. Then, the empirical relationships are investigated using ordinary least squares (OLS), random effects (RE), fixed effects (FE) and instrumental variables (IV) models. The final section explores whether there are income and gender disparities in the magnitude of the measured effects.

2. Major policies of the 1990s: the one child family policy and health sector reform

2.1 The one child family policy

Introduced formally in January 1979, the one child family policy refers not to one policy, but to a cornucopia of legislation and policy prescriptions designed to reduce fertility. Over time, benefits associated with compliance have included extra food rations, better housing, health subsidies, and allotments of farmland. Punitive measures have included fines, the loss of parents' jobs and the denial of workplace promotions or privileges. At the policy's peak, second births were forbidden altogether, except in extraordinary circumstances.

It is not easy to generalize about the characteristics of the one child policy, for at least two reasons: First, throughout its implementation, the one child policy has been a very much decentralized policy, resulting in substantial geographic variation in the policy's rules and the stringency of its implementation – even across very small administrative areas (Scharping 2003). Second, the severity of the policy has varied substantially over time. From an incentive-based structure at its introduction, the policy tightened increasingly until second births were effectively forbidden. In 1983, however, coercive enforcement was prohibited and the range of conditions under which couples could have two children started to expand (Li

2004). In 1988, about 12% of the Chinese population lived in areas with universal permission to have a second child and a further 50% lived in provinces extending a second-child permit to peasant households under certain conditions, and by 2001 about 60% of the Chinese peasant population and about 5% of the urban population was eligible for second-child permits (Scharping 2003).

2.2 Health sector reform

The second area of policy on which this paper focuses is a protracted period of health system reform which appears to have had important consequences for at least three dimensions of access to care, namely the price of healthcare, the quality of services, and the service mix.

In the public sector, there was evidence of rising user fees and out-of-pocket expenditures through the 1990s (World Bank 1997), possibly attributable to the introduction of hard budget ceilings (Hsiao 1995; Liu and Mills 2002), as well as declining insurance coverage and higher co-payments (Henderson et al. 1995; Akin, Dow and Lance 2004). In addition, government spending on preventive facilities (Liu 2004) and lower levels of service provision (Liu and Mills 2002) fell substantially. It is argued that these changes in health system financing had deleterious consequences for child health and nutritional status by increasing the delay in seeking treatment (Liu and Mills 2002), reducing the utilization of pre- and post-natal care (Anson 2004), and contributing to the fall in immunization coverage (Liu and Mills 2002). In addition, the emergence of unregulated private healthcare providers – another key feature of the reform – resulted in tremendous variation in the quality of care that is provided (Lim et al. 2004).

2.3 Descriptive evidence of policy change from the CHNS

The policy changes described in sections 2.1 and 2.2 are supported by data from the CHNS:

One plausible measure of the effect of the one child policy is the change in the percentage of only-children, i.e. the percentage of children under 12 years of age without siblings of (under the age of 18). The CHNS data show that, in the seven provinces considered in this study, there has been a sharp increase in the number of only-children under 12. In 2000, 42.8% of children were only-children compared to only 27.1% of children in 1991, equivalent to a 57.9% increase. However, not all spatial and temporal differences in the percentage of only-children can necessarily be attributed to the one child policy, which is why this paper does not claim to estimate the total effect of the one child policy on nutritional status, but rather the effect of being an only-child.

The description of trends in healthcare access and quality is also supported by the data. Financial access to healthcare facilities appears to have deteriorated between 1991 and 2000: the mean (real) price of a cold or influenza-related visit more than doubled from 3.8 yuan to 8.8 yuan. In addition, the full economic cost of accessing services increased: travel costs increased significantly from a mean of 0.05 real yuan in 1991 to 0.43 real yuan in 2000 – although rising costs may reflect increased use of motorized transport, rather than bicycles or walking. Travel time to the closest healthcare facility did not fall significantly between 1991 and 2000. On the other hand, to the extent that mean waiting time and the availability of medicines can be used as indicators of the quality of treatment available, there appear to have been some improvements. The mean waiting time to see a healthcare worker halved over the decade from a mean of 18 minutes down to 9 minutes, and while there was little change between 1991 and 1997, by 2000 there had been a statistically significant increase in the availability of medicine at facilities.

Table 1: Trends in accessibility and quality of health facilities in seven provinces, 1991 to 2000

	1991	1993	1997	2000
For the closest health facility:				
Mean travel time (minutes)	10	9.7	9.8	10.7
Mean real cost of travel (real yuan)	0.03	0.06	0.07	0.13
Mean waiting time (minutes)	17.6	15.2	11.7	8.7
Mean real cost of treating a cold	3.8	5.6	6.2	8.8
% households reporting medicine usually available	91.1%	93.3%	92.8%	93.6%

Source. China Health and Nutrition Survey, 1991-2000.

Note. Information was collected on every facility that the household can use, and the means are calculated for the facility that is closest (in distance) to the household. Values are calculated for households with at least one child under 12. F tests show that the differences between mean values for 1991 and 2000 are all statistically significant at the 1% level, except for the "percentage reporting medicine generally available" which is significant at the 10% level.

3. The theoretical model

The theoretical approach adopted in this paper has its origins in Becker's microeconomic model of household production (Becker 1965; Becker 1981) and the expansion of this model to the demand for health (Grossman 1972). The application of this model to child health status is well-documented (Behrman and Deolalikar 1988; Strauss and Thomas 1995; Currie 2000).

It assumes a one-period unitary model of decision-making whereby a household maximizes a **utility function** that depends on the consumption of commodities including a composite (market-purchased) household consumption good (X_{jt}), the consumption of leisure by all household members (L_{jt}) and child health or nutritional status (H_{ijt}), conditional on a set of household taste and preferences shifters (W_{jt}):

$$U_{jt} = U(X_{jt}, L_{jt}, H_{ijt}; W_{jt}) \qquad (1)$$

The household faces three constraints:

(i) a health production function representing the technology available to the household to transform available inputs into the health of the child. According to this function, the health or nutritional status of child *i* in household *j* in period *t* is:

$$\mathrm{H}_{\mathrm{ijt}} = h(M_{ijt}, Z_{ijt}, V_{ijt}; G_{ijt}, F_{jt}) \qquad (2)$$

where nutritional status depends on M_{ijt}, the endogenous material health inputs (such as immunizations, nutritional intake, healthcare utilization); on Z_{ijt}, the exogenous characteristics of the child and the household in which it lives; and on $\mathrm{V}_{\mathrm{ijt}}$, the endogenous child health-related time inputs. Child health is also conditional on two vectors of exogenous child-specific (G_{ijt}) and household/community-specific (F_{jt}) variables that are time-varying or time-invariant indicators of health technology[1]. As such, they influence the choice of health inputs and/or the efficiency with which existing inputs are combined to produce child health, and in so doing affect the shape of the health production function. Exogenous child-specific factors (G_{ijt}) could include age, sex and whether the child is an only-child. An example of an exogenous endowment at the household-level (F_{jt}) is the education of the caregiver (typically the mother) who makes health-related decisions that impact on the child's health. At the community-level, prevailing cultural norms regarding child health practices, the availability of healthcare facilities, relative food prices and communication or transportation infrastructure are examples of F_{jt} variables.

(ii) a **time endowment** reflecting the total time available to the household (N_{jt}) to allocate between wage labor (T_{jt}), leisure (L_{jt}) and health-related activities (V_{jt})

$$N_{jt} = T_{jt} + L_{jt} + V_{jt} \tag{3}$$

(iii) the **budget constraint** representing the total financial resources, consisting of wage income ($w_t T_{jt}$), non-wage income (Y_{jt}) and net assets (A_{jt}), available to the household with which to purchase market-produced health-related inputs (M_{ijt}) and other consumption goods (X_{jt}) at prices p^M and p^X respectively

$$p_t^X X_{jt} + p_t^M M_{ijt} = w_t T_{jt} + Y_{jt} + A_{jt} \tag{4}$$

Solution of the household's optimization problem yields the reduced-form health demand equation:

$$\mathrm{H}_{\mathrm{ijt}} = h(m(p_t^X, p_t^M, w_t, N_{jt}, Y_{jt}, A_{jt}; G_{ijt}, F_{jt}), Z_{ijt}, v(p_t^X, p_t^M, w_t, N_{jt}, Y_{jt}, A_{jt}; G_{ijt}, F_{jt}); G_{ijt}, F_{jt}) \tag{5}$$

The particular functional form $h(\;)$ depends on the underlying functions characterizing household preferences and health production.

The key programmatic variables of interest enter the model as follows:

Being an only-child is expected to affect child nutritional status in several ways. First, having one child rather than many children increases the *per capita* availability of household resources that can be directed towards enhancing child nutritional status. Second, it may also increase the *total* household resource availability since women who have fewer children tend to spend more time in wage-employment (Entwisle and Chen 2002). The latter effect may be partially captured by the income variable. Third, the incentives and sanctions associated with the one child policy have implications for child health and nutritional status: some may act through an income effect (e.g. fines may reduce money available for child-raising or grants for single children may increase it)[2] and others through a price effect (e.g. where healthcare subsidies are provided for only-children). Fourth, to the extent that the one child policy may alter prenatal and obstetric care-seeking behavior, it holds implications for birth outcomes which, in turn, are correlated with nutritional status later in childhood. The potential direction of this effect is ambiguous: Doherty, Norton and Veney (2001) show that the one child policy has been a significant deterrent to the utilization of prenatal and obstetric care among women with unapproved pregnancies, while Festini and de Martino (2004) have argued that the close monitoring of compliance with some of the terms of the one child policy, such as compulsory contraception, may have brought more women into contact with healthcare services than would otherwise have been the case. For the last two reasons, it is expected that the magnitude of the effect on nutritional status of being an only-child will be greater in China than is typically found in other countries where the effect of being an only-child operates only through the first two mechanisms.

Healthcare is an endogenous health input (M_{ijt}). The dominant pathway by which access to quality preventive and curative healthcare services is expected to influence nutritional status is by reducing the incidence, severity or duration of childhood disease. Children, and especially young children, who fall ill may rapidly deplete their nutritional stores because of reduced intake of food, poor absorption of nutrients and the increased demands of combating disease (Scrimshaw and SanGiovanni 1997; Allen and Gillespie 2001), leading to growth retardation. A second theoretical pathway is via the provision of appropriate antenatal care and advice to pregnant women which may improve birth outcomes and, therefore, the nutritional status of children later in life (Schmidt et al. 2002). Indeed, recent studies have found a positive effect of access to health services on child nutritional status in Vietnam (Glewwe, Koch and Nguyen 2002) and in Colombia and Peru (Behrman and Skoufias 2004).

4. Analytical approach

4.1 Data

Data are drawn from a large-scale panel survey, the China Health and Nutritional Survey (CHNS). This paper utilizes four waves, from 1991 through 2000, and includes the seven provinces of Jiangsu, Shandong, Henan, Hubei, Hunan, Guangxi and Guizhou. The 1989 wave is excluded because in that year anthropometric data were not collected for children over the age of six years.

In each of the provinces, a multistage, random cluster procedure was used to draw the sample (see Henderson et al. 1994, for example, for an elaboration of the sample design). Counties in each province were stratified by income (low, middle, and high) and a weighted sampling scheme was used to randomly select 4 counties in each province. In addition, the provincial capital and a lower income city were selected where feasible. Villages and townships within the counties, and urban and suburban neighborhoods within the cities, were selected randomly.

In CHNS 1991, only individuals belonging to the original 1989 sample were re-interviewed; in CHNS 1993, all the new households that had been formed from the 1991 sample households who resided in sample areas were added; in CHNS 1997, all newly-formed households who resided in sample areas were once again added, plus additional households and communities to replace those no longer participating; and in CHNS 2000, newly-formed households, replacement households, and replacement communities were again added. In the seven provinces included in this analysis, there were 3,207 households in 1991, 2,986 households in 1993, 3,296 households in 1997 and 3,237 households in 2000 across 176 primary sampling units (or communities); there were 13,229 individuals in 1991, 12,492 individuals in 1993, 14,268 individuals in 1997 and 13,657 individuals in 2000.

4.2 Measurement of the variables

The dimension of nutritional status with which this paper is concerned is linear growth – a measure of protein-energy malnutrition. Consequently, the dependant variable used is the height-for-age z-score (or the recumbent length-for-age z-score in the case of children under the age of 36 months). This measure is constructed on the basis of the 2000 Center for Disease Control (CDC) growth charts (Kuczmarski 2002). In line with the recommended exclusion ranges for implausible values, children with height-for-age z-scores that are less than -5 and greater than +3 have been excluded from the analysis (WHO 1995).

The measure of income used is the log of real per adult income. This measure is carefully constructed from the detailed income data in the CHNS to include multiple net cash income streams, the market value (as estimated by the household head) of in-kind income and the (estimated) market value of the quantity of home production consumed. Income is adjusted for temporal and spatial price variation using the China State Statistical Bureau's (SSB) annual consumer price index ratio. Since fertility may be endogenous (Schultz 1976; Wolfe and Behrman 1992), real household income is deflated by the total number of adults rather than total household size.

An only-child is defined as someone under the age of 18 without siblings that are younger than 18.

Four variables capturing physical and financial access to healthcare are constructed. These variables are the non-self cluster means of one-way travel time (in minutes), travel cost (in real yuan), the cost of treating a common cold (in real yuan) and waiting times at the healthcare facility that is closest to the household facility. In addition, a measure of medicine availability (where the availability indicator = 1 if at least 90% of respondents in the community cluster report regular availability) is included to capture healthcare quality. The use of cluster (i.e. community) means of the healthcare facilities helps to reduce the potential correlation with unobserved household characteristics through simultaneity. The use of *non-self* cluster means allows the introduction of a little household-level variation into these variables since each observation within a cluster will take on a different value, thus avoiding potential collinearity in the community-level fixed effects models. Missing values are imputed for those households with missing values on one or more healthcare characteristics by assigning the cluster/community mean value to these households.

Other variables relevant to the production of child nutritional status are included in the model as covariates. Height-for-age z-scores are modeled as a quadratic function of the child's age (in years) to capture the typical growth-faltering pattern of infants in developing countries; sex is a binary variable; mother's educational attainment is included as a vector of four dummy variables indicating significant thresholds of scholastic achievement, namely no formal education, 1 to 3 years of primary school, 4 to 6 years of primary school, 1-3 years of lower middle school, and 1-3 years of upper middle school or more; mother's height, measured in centimeters, is included to capture the child's unobserved growth potential or genetic endowment.

4.3 Model specification

In the empirical formulation of the theoretical model outlined in section 3, the nutritional status of child *i* in community *j* at time *t*, (H_{ijt}), can be modeled as a function of *k* regressors that may be time-varying – such as the child's age, income, whether the index child is an only-child and access to health services – or time-invariant – such as sex and genetic endowment.

This paper assumes that the composite error term ε_{ijt} has a random component μ_{it} – that varies across children and over time – and a fixed component v_j – that does not vary across children within specific communities over time, but may vary across communities in the sample. Year dummies capture any trending in nutritional status over time, i.e. aggregate time effects that have the same influence on the nutritional status of all children regardless of their other observed and unobserved characteristics.

$$H_{ijt} = \beta_0 + \beta_1 X_{ijt1} + ... + \beta_k X_{ijtk} + \delta_1 1993 + \delta_2 1997 + \delta_3 2000 + \varepsilon_{ijt} \, where \, \varepsilon_{ijt} = \mu_{it} + \nu_j \qquad (6)$$

The correct specification of a reduced-form child health demand model would not include a measure of household income. So, strictly speaking, this is a conditional demand model since nutritional status is modeled as a function of some inputs from the structural health production function and some variables from the reduced-form relations. Although this model cannot reveal all of the structural parameters nor the pathways through which particular policies may operate, it is informative about the effect of income and allows the direct effect on health of other variables, such as education, to be separated from any indirect effect that might operate through income.

The foundation of this analysis is a set of pooled OLS models that examine the relationship between nutritional status, on the one hand, and income and policy variables, on the other hand. Observations on all children who were under the age of 12 in each survey year are pooled and the panel is treated as if in each wave a new random sample was drawn from the population, thus exploiting only between-group variation. Because preliminary models exhibit evidence of heteroskedasticity (according to the Breusch-Pagan test), because observations within community clusters are not independent and because (in panel data) there is likely to be serial correlation within communities and/or across individual observations over time, Huber community-clustered robust standard errors are used to avoid underestimating the true standard errors.

Second, a random effects (RE) model is estimated which, if all its assumptions hold, is a more efficient estimator than a pooled OLS model.

Third, since unobserved (time-invariant) community factors may not be uncorrelated with all the explanatory variables, a key assumption of the RE model, a community fixed effects (FE) model is estimated. An example of such a time-invariant factor could be the general level of economic development of the community (relative to other communities) which could exert a direct effect on nutritional status, but also be correlated with adult income, the number of children and the characteristics of healthcare facilities. In particular, the risk of bias posed by the potential endogeneity of the placement and characteristics of healthcare services is problematic. If the placement and/or characteristics of healthcare facilities are related to other unobserved factors that could also be correlated with nutritional status, such as population density, level of economic development and administrative rank, the coefficients of the healthcare variables may be biased. To the extent that these unobserved factors are time-invariant, though, their influence can be negated when a community fixed effects model is used.

Fourth, in an attempt to overcome the potential endogeneity of the income variable, an instrumental variables (IV) approach to the OLS and FE models is also explored.

5. Results

Initial OLS models support the foundational hypothesis of this paper. Real adult income is found to be a significant correlate of improved child nutritional status in a basic model that includes the income variable and other covariates, but excludes the policy variables (see Table 2, column I). When policy variables are introduced, the effect of income remains significant, but its effect is substantially mitigated (see Table 2, columns II and III). The income coefficient of 0.1 in the basic model falls to 0.08 when being an only-child is controlled for, and falls still further to 0.06 when healthcare variables are introduced.

5.1 Specification tests for the OLS, RE and FE models

The full pooled OLS model is rejected in favor of an RE model with community-level random effects. The RE model (see Table 2, column IV) produces smaller coefficients than the OLS model on the income and only-child variables and the results of the Breusch-Pagan Lagrangian multiplier test [chi2(1)=872.06] indicate that these differences are significant, meaning that the unobserved effects are relatively important. However, the RE model suffers from heteroskedasticity (according to the Breusch-Pagan test) and serial error correlation (according to the AR(1) test for autocorrelation), and the RE stimates are likely to be inconsistent.

Next, the RE model is compared to a community-level fixed effects model (see Table 2, column V). The Hausman test [chi2(18)=118.46] confirms that the FE model is the preferred model and that, with these data, the RE model produced biased estimates. This implies that the FE model is also preferred to the pooled OLS model since the RE model is more efficient than the pooled OLS model. Ex-post Huber standard errors are used because, although in FE models the time-invariant error component no longer causes the error terms to be clustered, errors may still be correlated if there are omitted community-specific time trends. Fixed effects are strongly and significantly positive [F(175, 6005)=5.07].

5.2 Findings of OLS, RE and FE models

Being an only-child is found to be one of the most important predictors of nutritional status: on average, and controlling for other factors, only-children have height-for-age z-scores that are greater than those of children with siblings. The coefficient remains highly significant across all specifications, but is approximately half the size in the FE model (0.10) than it was in the RE model (0.21) and one-third of the size that it was in the OLS model (0.33) (see Table 2). These differences suggest that the exclusion of community fixed effects did indeed bias the coefficients in the predicted direction. With respect to the healthcare variables, in the preferred FE model there is no evidence of a significant relationship between the access to and quality of care and improvements in nutritional status.

In all models, the magnitude and direction of the coefficients of most of the other variables conform to expectations. The coefficients on the sex variable show that boys' height-for-age z-scores are not significantly higher than girls'. Height-for-age z-scores deteriorate significantly as children age, but at a decreasing rate, indicating substantial age-related growth failure. Maternal education is significantly associated with favorable anthropometric outcomes, but only at the higher levels of educational attainment, specifically beyond primary school. Mother's height is also highly significant, with the smallest standard error of all variables.

The income variable, while highly significant in the OLS and RE models, becomes insignificant once community fixed effects are controlled for. This effect will be further explored in the next section, where income will be instrumented.

The positive coefficients on all three of the time dummies provide clear evidence of secular improvements in nutritional status in the first part of the decade, but with less change towards the end of the decade. Controlling for changes in other time-varying factors, average height-for-age z-scores increased by more than one-tenth of a standard deviation between 1991 and 1993 and by more than a quarter of a standard deviation between 1991 and 1997. Thereafter, however, there appears to have been no further significant improvements in nutritional status: the null hypothesis that the 1997 and 2000 dummies are equal could not be rejected by F-tests.

5.3 Findings of IV models

The income variable is potentially endogenous due to the joint determination of income and child nutritional status, through the labor/time allocation decision, in the utility maximization problem. In other words, the labor/time allocation decision affects both income, via its effect on (especially female) labor supply, *and* child nutritional status, through influencing the time available for child care activities. Failure to control for this simultaneity will generate biased estimates of income, and also of other covariates that are correlated with income. Another potential source of endogeneity is measurement error in income. While simultaneity will tend to bias OLS estimates upwards, measurement error will tend to bias the estimates towards zero. This may partially explain both the large income coefficient in the OLS model and the failure to find a significant effect on the income variable in the FE model. Consequently, an instrumental variable approach is explored.

Finding instruments that are simultaneously strong predictors of income in the first stage regression and also satisfy the exclusion criteria (i.e. are uncorrelated with the error term of the main equation) is challenging. There is also a trade-off between the consistency of IV estimates and the efficiency of OLS estimates: while an IV approach reduces the problems of bias introduced by endogeneity, it causes the variance of the parameters to increase. After giving consideration to a number of different instruments that have been used in previous studies of child nutritional status[3], two instruments – housing floorspace (measured in square meters) and the number of salary earners in the household – were selected. In the instrumental variables OLS model (see Table 2, Column V), these instruments were found to (i) be individually and jointly significant (F-test of joint significance equivalent to F(2,156)=60.28) in the first stage regression; (ii) together with other covariates, explain a sufficient share of the variation in income in the first stage regression (R-squared of 0.165); and (iii) be validly excluded from the main nutrition equation in the instrumental variables model. With respect to the exclusion criteria, the Hansen J test of overidentification could not be rejected (P>chi2=0.197) and regressions of the residuals from the main equation on the instruments and other covariates did not yield significant coefficients on the instruments, suggesting that the instruments are validly excluded from the main equation.

The Hausman test recommends rejecting the pooled OLS estimates in favor of the IV estimates, confirming that income was indeed endogenous in the OLS model. In the IV model, the income coefficient is significant and very large (0.46) – nearly seven times as large as the income coefficient in the original pooled OLS model – supporting the hypothesis that there was substantial underestimation of the income coefficient due to measurement error. The only-child variable, waiting times at health facilities and travel times to health facilities are significant and have magnitudes that are very similar to the original OLS model.

Once community-level fixed effects are controlled for in the IV model[4], however, the significance of (instrumented) income disappears, as do the effects of most of the healthcare variables. The only-child variable remains significant, but the magnitude of the effect has fallen by almost one-third compared to the instrumented OLS model. Being an only-child increases height-for-age z-scores by 0.12 of a standard deviation. The magnitude of this effect is notably larger than then the FE model without instrumental variables where the coefficient on the one child variable was 0.1. Waiting time at the healthcare facility is the only measure of healthcare that is significant, although only at the 10% level and in an unanticipated direction. One possible explanation for this result is that waiting time is associated with good quality care – because households are willing to wait longer at facilities where they can receive better quality care – which is, in turn, associated with improved nutritional status.

Table 2: Regression of height-for-age z-scores in pooled OLS, RE, FE and IV models

	I OLS with income	II OLS with one-child	III Full OLS	III RE	IV FE	V IV OLS	VI IV FE
Sex (Male=1)	-0.021	-0.031	-0.029	-0.012	-0.002	-0.02	0.001
	(0.039)	(0.038)	(0.036)	(0.026)	(0.033)	(0.036)	(0.026)
Age (years)	-0.085	-0.055	-0.059	-0.068	-0.075	-0.083	-0.081
	(0.021)***	(0.021)***	(0.021)***	(0.019)***	(0.021)***	(0.021)***	(0.020)***
Age squared	0.006	0.004	0.004	0.005	0.005	0.005	0.005
	(0.001)***	(0.001)***	(0.001)***	(0.001)***	(0.001)***	(0.001)***	(0.001)***
Maternal education (“no formal education” is omitted):							
1-3 yrs primary	0.005	-0.021	-0.012	0.017	0.041	-0.05	0.044
	(0.102)	(0.1)	(0.095)	(0.062)	(0.088)	(0.098)	(0.063)
4-6 yrs primary	0.059	0.038	0.025	0.024	0.023	-0.033	0.021
	(0.089)	(0.087)	(0.082)	(0.047)	(0.076)	(0.088)	(0.048)
1-3 yrs lower middle	0.298	0.245	0.201	0.152	0.118	0.081	0.105
	(0.081)***	(0.078)***	(0.074)***	(0.044)***	(0.066)*	(0.083)	(0.047)**
1-3 yrs upper or more	0.523	0.416	0.336	0.251	0.173	0.188	0.159
	(0.098)***	(0.091)***	(0.085)***	(0.053)***	(0.081)**	(0.093)**	(0.057)***
Mother’s height (cm)	0.063	0.06	0.056	0.05	0.047	0.054	0.046
	(0.005)***	(0.005)***	(0.005)***	(0.003)***	(0.004)***	(0.004)***	(0.003)***
Log real per adult income (yuan)	0.1	0.084	0.064	0.045	0.021	0.462	0.138
	(0.026)***	(0.024)***	(0.024)***	(0.019)**	(0.022)	(0.087)***	(0.087)
Year dummies:							
1993	0.142	0.143	0.12	0.127	0.148	0.137	0.149
	(0.026)***	(0.025)***	(0.033)***	(0.033)***	(0.027)***	(0.037)***	(0.034)***
1997	0.279	0.256	0.271	0.288	0.314	0.281	0.31
	(0.046)***	(0.044)***	(0.051)***	(0.039)***	(0.046)***	(0.051)***	(0.040)***
2000	0.317	0.24	0.27	0.296	0.34	0.377	0.357
	(0.063)***	(0.060)***	(0.073)***	(0.050)***	(0.063)***	(0.070)***	(0.054)***
Index child is only-child		0.426	0.325	0.208	0.104	0.311	0.119
		(0.051)***	(0.047)***	(0.035)***	(0.046)**	(0.049)***	(0.039)***
Non-self cluster means of healthcare variables:							
Cost of treating cold (real yuan)			0.015	0.014	0.007	0.005	0.003

			(0.005)***	(0.004)***	(0.005)	(0.005)	(0.003)
Travel time (minutes)			-0.017	-0.005	0.003	-0.016	0.006
			(0.004)***	(0.003)**	(0.003)	(0.004)***	(0.005)
Cost of travel (real yuan)			0.122	0.166	0.158	0.133	0.161
			(0.107)	(0.083)**	(0.122)	(0.133)	(0.088)*
Waiting time (minutes)			0.007	0.006	0.002	0.006	0.003
			(0.002)***	(0.001)***	(0.002)	(0.002)***	(0.002)
Medicine regularly available			-0.032	0.032	0.064	-0.018	0.055
			(0.052)	(0.036)	(0.042)	(0.053)	(0.039)
Constant	-11.816	-11.329	-10.513	-9.533	-8.771	-12.895	-9.519
	(0.779)***	(0.739)***	(0.692)***	(0.436)***	(0.607)***	(0.796)***	(0.707)***
Observations	6249	6249	6199	6199	6199	6199	6199
Robust standard errors, clustered at the community level, in parentheses ; 176 community fixed effects							

Source. China Health and Nutrition Survey, 1991-2000.
Note. * significant at 10%; ** significant at 5%; *** significant at 1%.

5.4 Equity considerations: are there income and gender disparities in the effects of the policy variables?

To explore whether the impact on nutritional status of being an only-child and having access to healthcare is more salient for children with certain characteristics, and in particular for children that belong to relatively disadvantaged subgroups, interaction terms are introduced into the models.

There do not appear to be any income or gender disparities in the effect of the only-child variable. While being an only-child was a highly significant variable in all models estimated in this paper, the nutritional status of only-children is not significantly better in higher income households. Also, boys who are only-children do not have z-scores that are significantly different from girls who are only-children.

Table 3: Regressions of height-for-age z-scores on only-child interaction terms

	Gender interaction		Income interaction	
	OLS	FE	OLS	FE
Sex (Male=1)	-0.024	-0.002	-0.029	-0.002
	(0.042)	(0.037)	(0.036)	(0.033)
Log real per adult income (yuan)	0.064	0.021	0.052	0.024
	(0.024)***	(0.022)	(0.027)*	(0.024)
Index child is only-child	0.336	0.102	-0.048	0.206
	(0.061)***	(0.061)*	(0.367)	(0.348)
Interaction terms:				
Sex*Only-child	-0.019	0.003		
	(0.073)	(0.072)		
Income*Only-child			0.051	-0.014
			(0.05)	(0.048)
Observations	6199	6199	6199	6199
No. of community fixed effects		176		176
Robust standard errors, clustered at the community level, in parentheses				

Source. China Health and Nutrition Survey, 1991-2000.
Note. Only select coefficients shown; other variables are the same as in the full pooled OLS and FE models (see Table 2).
*significant at 10%; **significant at 5%; ***significant at 1%.

The interaction of healthcare variables and income also produces insignificant coefficients (see Table A1 in the Appendix). This implies that the extent to which the availability, quality and cost of healthcare will affect child nutritional status does not depend on the amount of income that the household has available to purchase that care. Distance to the healthcare facility (as measured by traveling time) is the only healthcare variable for which a statistically significant gender effect is observed (see Table 4). Since longer traveling times are associated with poorer nutritional status (i.e. lower height-for-age z-scores), the positive coefficient on the interaction term indicates that the association is somewhat weaker for boys than for girls.

Table 4: Regressions of height-for-age z-scores on interactions between gender and healthcare

	Treatment cost		Travel time		Travel cost		Waiting time		Medicine	
	OLS	FE	OLS	FE	OLS	FE	OLS	FE	OLS	FE
Sex (Male=1)	-0.051	-0.018	-0.115	-0.09	-0.029	-0.009	-0.035	-0.016	0.046	0.054
	(0.048)	(0.044)	(0.057)**	(0.054)*	(0.037)	(0.034)	(0.051)	(0.047)	(0.059)	(0.059)
Non-self cluster means of healthcare variables:										
Cost of treating cold (real yuan)	0.013	0.005	0.015	0.007	0.015	0.007	0.015	0.007	0.015	0.007
	(0.006)**	(0.005)	(0.005)***	(0.005)	(0.005)***	(0.005)	(0.005)***	(0.005)	(0.005)***	(0.005)
Travel time (minutes)	-0.017	0.003	-0.021	-0.001	-0.017	0.003	-0.017	0.003	-0.017	0.003
	(0.004)***	(0.003)	(0.004)***	(0.004)	(0.004)***	(0.003)	(0.004)***	(0.003)	(0.004)***	(0.003)
Cost of travel (real yuan)	0.121	0.158	0.116	0.154	0.123	0.097	0.122	0.159	0.122	0.158
	(0.107)	(0.122)	(0.104)	(0.116)	(0.119)	(0.134)	(0.107)	(0.122)	(0.107)	(0.122)
Waiting time (minutes)	0.007	0.002	0.007	0.002	0.007	0.002	0.007	0.002	0.007	0.002
	(0.002)***	(0.002)	(0.002)***	(0.002)	(0.002)***	(0.002)	(0.002)***	(0.002)	(0.002)***	(0.002)
Medicine regularly available	-0.031	0.064	-0.032	0.063	-0.032	0.064	-0.032	0.064	0.021	0.102
	(0.052)	(0.043)	(0.052)	(0.042)	(0.052)	(0.042)	(0.052)	(0.042)	(0.059)	(0.054)*
Interaction terms:										
Sex*Cost	0.004	0.003								
	(0.006)	(0.005)								
Sex*Travel time			0.008	0.009						
			(0.004)**	(0.003)**						
Sex*Travel cost					-0.002	0.129				
					(0.119)	(0.119)				
Sex*Waiting time							0	0.001		
							(0.002)	(0.002)		
Sex*Medicine									-0.098	-0.072
									(0.060)	(0.060)
Observations	6199	6199	6199	6199	6199	6199	6199	6199	6199	6199

Robust standard errors, clustered at the community level, in parentheses;

Source. China Health and Nutrition Survey, 1991-2000.
Note. Only select coefficients shown; other coefficients are the same as in the full pooled OLS and FE models (see Table 2).
*significant at 10%; **significant at 5%; ***significant at 1%.

6. Discussion and policy implications

6.1 Socioeconomic determinants

This analysis yields a number of findings that are broadly consistent with the literature on the socioeconomic determinants of child nutritional status, but with some variation that may be specific to the Chinese context. There is evidence of substantial age-related growth failure, as is observed throughout the developing world (Shrimpton et al. 2001), and which is typically caused by factors such as premature cessation of exclusive breastfeeding, the failure to timeously introduce appropriate complementary foods of sufficient quality and quantity and exposure to disease (Allen and Gillespie 2001). Boys' z-scores are not significantly different from girls', implying that any differences in growth attainment are fully attributable to biological differences between the sexes and, thus, captured by the use of z-scores. Mother's height is indeed found to be an important determinant of a child's linear growth, either because it captures genetic endowment (as is commonly assumed) or because of its effect on children's birth weight (Xu et al.; Blumenfield et al. 2006). As has been found in analyses of other lower- and middle-income countries (see, for example, Glewwe 1999; Handa 1999; Thomas, Strauss and Henriques 1991), this analysis produces particularly large and favorable effects of maternal education, but this is observed only at higher levels of educational attainment. This suggests that in China, as Christiaensen and Alderman (2004) found in Ethiopia, it is more than just basic information-processing skills or functional literacy (i.e. the competencies associated with early education) that are critically important for child health. It is also found that, compared to the initial model without only-child and healthcare variables (see Table 2, column I), the inclusion of the only-child variable weakens the coefficient on maternal education. This result could be explained by a direct effect of education on the fertility decision – highly educated mothers tend to have fewer and healthier children (Schultz 1976) – and/or a possible spurious correlation between maternal education and number of children since the one child policy is more strictly enforced in urban areas (Scharping 2003) where educational opportunities are also better.

With respect to income, while it is initially surprising that the preferred FE model does not yield a significant coefficient on the income variable, this result should not be interpreted as evidence that income does not affect nutritional status or that policies aimed at improving household incomes would not be effective in improving child nutritional status. Since the introduction of community fixed effects absorbs between-community variation, this result may simply imply that most of the relationship between income and nutritional status observed in the initial OLS and RE models is attributable to variation in income between communities rather than between households within communities.

6.2 Policy variables

A specific objective of this paper was to consider the effects of specific government policies on nutritional status.

Access to healthcare was found to be an important correlate of nutritional status, at least in the OLS models. An increase in the time taken to travel to healthcare facilities was associated with worse nutritional outcomes and, if the cost of treatment and waiting time can be assumed to be indicators of facility quality, then better facility quality was associated with better nutritional status. The financial cost of care was not significant, however, suggesting that the increasing user fees and decreasing insurance coverage associated with the healthcare transition may not have had too detrimental an effect on child nutritional status. One should be careful not to overstate the importance of healthcare characteristics in these models, though, because once time-invariant intra-community variation is removed in an FE model, the effects of the healthcare variables are no longer significant.

Further, while the results of this analysis should not be interpreted as a measure of the effect of the one child policy *per se* (since the variable used in the analysis only captures one outcome of the policy), the analysis is illuminating with respect to the nutritional consequences of reduced fertility. Controlling for the socioeconomic characteristics of the household, only-children had height-for-age z-scores that were a third (OLS models) to a fifth (FE models) of a standard deviation larger than that of children with siblings. The results are consistent with a conclusion that any policy measure that can effectively encourage parents to have only one child would have large positive consequences for child nutritional status. Moreover, it is likely that the difference in nutritional status between only-children and children with siblings in China is larger than one would find in other countries where reduced fertility is the result of other policies, such as the promotion of family-planning methods and the enhancement of female autonomy. This is because, as pointed out in Section 3, the particular rewards and sanctions associated with the one child policy have direct consequences for child health. For example, only-children often received health subsidies and other health-promoting inputs that were denied to subsequent children, and women may have failed to seek healthcare for unapproved higher order births out of fear of financial penalties.

Finally, it is interesting to find that the effects of these key policy-related variables are largely income- and gender-neutral. Higher incomes do not significantly change the effect on nutritional status of access to care and it would appear that only-children are better off than other children, regardless of household economic status. With respect to gender disparities, there is no evidence that boys who are only-children have z-scores that are significantly different from girls who are only-children. However the association between one of the healthcare variables, namely distance to healthcare facilities, is weaker for boys than for girls. This could be interpreted as suggesting that the distance to health facilities is less of an obstacle to seeking care when the child is a boy, implying a weak son-preference.

Endnotes

[1]F_{ijt} and G_{ijt} variables can be thought of as exogenous productivity shifters, which may be either permanent or temporary shocks (Currie 2000).

[2]The income effect of fines or subsidies will be captured by the "only-child" variable since this income source is not included in the constructed income measure.

[3]Other instruments that were explored include various combinations of the ownership or value of assets such as consumer durables, housing, equipment associated with household farms or enterprises, savings, land holdings and livestock ownership (see, for example, the wide variety used in the cross-country study of Haddad et al. 2003), community wages (see, for example, Attanasio, Syed and Vera-Hernandez 2004), and indicators of the industry or occupation of income earners (see, for example, Case, Lubotsky and Paxson 2002).

[4]In the fixed-effects model with instrumental variables, like the instrumented OLS model, the instruments are individually and jointly significant, explain a substantial share of the variation in income in the first stage regression and can be validly excluded from the main nutrition equation.

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Table A1: Regressions of height-for-age z-scores on interactions between income and healthcare

	Treatment cost		Travel time		Travel cost		Waiting time		Medicine	
	OLS	FE	OLS	FE	OLS	FE	OLS	FE	OLS	FE
Log real per adult income (yuan)	0.102	0.025	0.075	-0.007	0.07	0.028	0.095	0.031	0.047	0.007
	(0.034)***	(0.028)	(0.040)*	(0.034)	(0.026)***	(0.023)	(0.032)***	(0.026)	(0.044)	(0.042)
Non-self cluster means of healthcare variables:										
Cost of treating cold (real yuan)	0.073	0.012	0.015	0.007	0.016	0.007	0.016	0.007	0.015	0.007
	(0.037)*	-0.031	(0.005)***	-0.005	(0.005)***	-0.005	(0.005)***	-0.005	(0.005)***	-0.005
Travel time (minutes)	-0.017	0.003	-0.009	-0.016	-0.017	0.003	-0.017	0.003	-0.017	0.003
	(0.004)***	-0.003	-0.022	-0.02	(0.004)***	-0.003	(0.004)***	-0.003	(0.004)***	-0.003
Cost of travel (real yuan)	0.142	0.159	0.121	0.165	0.705	0.723	0.126	0.159	0.121	0.158
	-0.107	-0.122	-0.106	-0.126	-0.519	-0.612	-0.106	-0.122	-0.107	-0.122
Waiting time (minutes)	0.007	0.002	0.007	0.002	0.007	0.002	0.025	0.007	0.007	0.002
	(0.002)***	-0.002	(0.002)***	-0.002	(0.002)***	-0.002	(0.013)*	-0.012	(0.002)***	-0.002
Medicine regularly available	-0.032	0.064	-0.033	0.067	-0.032	0.064	-0.033	0.064	-0.197	-0.082
	-0.051	-0.042	-0.052	-0.042	-0.052	-0.042	-0.052	-0.042	-0.388	-0.347
Interaction terms:										
Income*Cost	-0.008	-0.001								
	(0.005)*	-0.004								
Income*Travel time			-0.001	0.003						
			-0.003	-0.003						
Income*Travel cost					-0.079	-0.077				
					-0.069	-0.075				
Income*Waiting time							-0.002	-0.001		
							-0.002	-0.002		
Income*Medicine									0.022	0.02
									-0.052	-0.046
Observations	6199	6199	6199	6199	6199	6199	6199	6199	6199	6199
Robust standard errors, clustered at the community level, in parentheses;										

Source. China Health and Nutrition Survey, 1991-2000.
Note. Only select coefficients shown; other coefficients are the same as in the full pooled OLS and FE models (see Table 2).
*significant at 10%; **significant at 5%; ***significant at 1%.